

CREDIT RATIONING AND MONETARY TRANSMISSION: EVIDENCE FOR PORTUGAL

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1 — Introduction

The purpose of this paper is to evaluate the existence of credit rationing and its relevance to monetary transmission in Portugal during the period from January 1990 to November 1997. Using monthly data we performed two types of tests: stickiness and causality tests. The paper is organised as follows: section two sets the theoretical and empirical background of credit rationing and its relevance to the monetary transmission mechanism; the stickiness tests are the object of section three; section four deals with the causality tests and section five is a conclusion.

2 — Credit rationing and monetary transmission

We define «credit rationing» as a situation where demand for loans exceeds supply at the prevailing interest rate. The price of the loan (the interest rate) does not fully adjust so that demand is not completely satisfied. The rationing of demand may be achieved in two ways: either borrowers do not receive the full amount of credit they have applied for (the so called «type I rationing») or some of the borrowers are simply turned down («type II rationing») ⁽¹⁾.

In a seminal paper, Stiglitz and Weiss (1981) argue that credit rationing may occur because of asymmetric information in credit markets. Banks (the lenders) have little information on the default risks of the applicants. They have therefore an incentive in not raising interest rates when demand exceeds supply. An adverse selection effect would occur following an interest rate increase that would drive away the less risky applicants. Moreover, an incentive effect would operate because borrowers would be tempted by riskier undertakings. In the end, an interest rate rise would not maximise bank profits as the number of defaults would increase.

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⁽¹⁾ This distinction was due to Keeton (1979) and is widely used in the credit rationing literature. Swank (1996) and Escária (1997) provide surveys on the theory of credit rationing.

Some authors have argued that rationing in credit markets could be of minor importance once contractual mechanisms like loan commitments and collateral are considered ⁽²⁾. The theoretical literature is therefore inconclusive and we, like Berger and Udell (1992), consider that the existence and importance of credit rationing remains an important empirical issue.

The possible existence of credit rationing opens the way to a credit channel in the transmission mechanism of monetary policy. The credit rationing hypothesis implies that monetary contraction would most probably lead banks to contract credit independently of the effect on loan demand of any increase in interest rates. Therefore, some degree of stickiness in the adjustment of credit rates to changes in money market rates is to be expected if credit rationing is to have any importance. Also, and still under the credit rationing hypothesis, one expects that monetary policy will have a direct effect on credit, independently of the interest rate channel. These two ideas underlie the two tests we describe in this paper.

We have included a brief review on the empirical evidence on credit rationing in a previous paper ⁽³⁾. Tests by surveyed authors basically rely on one of two ideas, also pursued here:

- i) If there is credit rationing, credit rates should not fully respond to changes in money rates (the so called «stickiness tests»). This is the avenue followed by Slovin and Sushka (1983) and Berger and Udell (1992);
- ii) If there is no credit rationing, the interest rate channel should prevail over the credit channel in the monetary transmission mechanism. This idea is usually exploited by means of «causality tests» and followed by King (1986) and Sofianos, Wachtel and Melnik (1990).

Results on the empirical relevance of credit rationing are different for the same economy (the US) according to period and performed test. Table 1, borrowed from Afonso and St. Aubyn (1997) summarises results obtained by the aforementioned authors:

⁽²⁾ See Besanko and Takor (1987) on collateral and Sofianos, Wachtel and Melnik (1990) on loan commitments.

⁽³⁾ See Afonso and St. Aubyn (1997).

TABLE 1

Some empirical evidence of credit rationing

| Author and date | Data frequency | Period and country | Test performed | Results | Evidence of credit rationing? |
|--------------------------------------|-----------------|------------------------------|---|------------|-------------------------------|
| Slovin and Sushka (1983) | Quarterly | 1952:3 to 1980:4 (USA) | Commercial loan rates adjust to market rates? | Yes | No |
| King (1986) | Quarterly | 1955:1 to 1979:3 (USA) | Loan rates are affected by bank liquidity? Loan supply responds to loan rate? | Yes Yes | Yes No |
| Sofianos, Wachtel and Melnik (1990). | Monthly | July 1973 to June 1987 (USA) | Direct causality from monetary policy to loans not under commitment? | Yes | Yes |
| | | | Direct causality from monetary policy to loans under commitment? | No | Yes |
| Berger and Udell (1992) | Quarterly | 1977:1 to 1988:2 (USA) | Commercial loan rate is «sticky» with respect to open-market rate? | Yes | Partially |
| | | | The commitment proportion of new loans rises when open-market rates are high? | No | No |

3 — Interest rate stickiness

The credit rationing hypothesis, namely as stated by Stiglitz and Weiss (1981), implies that credit rates do not adjust fully and/or quickly to changes in money market rates. If money market conditions worsen, banks will not tend to fully reflect these into higher interest rates for their customers.

In this section our empirical investigation is twofold. First, we test the hypothesis that money market rates movements are one to one reflected in credit market rates for private companies, no matter of how long this adjustment takes. Then, and having accepted the aforementioned hypothesis, we pursue to estimate the speed of adjustment ⁽⁴⁾.

Our data are monthly from December 1990 to November 1997. The money market rate is TBA ⁽⁵⁾. Credit market rates are the average credit rates offered by banks to private companies, according to credit length. Four lengths were considered: loans from 91 to 180 days, from 181 days to one year, from 2 to 5 years, and more than 5 years.

For several years, between 1978 up to 1990, Portugal displayed several forms of direct monetary control where quantity and price limits were enforced on the banking system and on the credit market. However, shortly after the 90s indirect monetary control started to be implemented and, in a period of more or less three years, credit and deposit rates were indeed believed to be the result of market demand and supply.

Figure 1 displays the TBA ⁽⁶⁾ and two credit rates to non-financial private enterprises after 1990. Visual inspection suggests that they are quite related. As one would surely expect, money market rates are always below credit rates, and falls in those rates seem to impinge on credit rates. Notice that there was a clear tendency for a fall in interest rates as Portuguese inflation felt dramatically in this period.

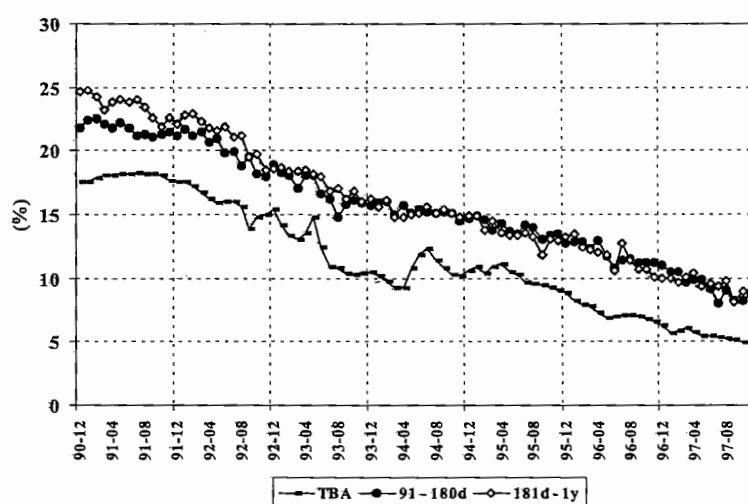
⁽⁴⁾ The empirical work in this section updates early empirical work by the same authors – see Afonso and St. Aubyn (1997).

⁽⁵⁾ Average rate of Treasury Bills auctions.

⁽⁶⁾ A similar relation can also be detected with for instance LISBOR6 (Lisbon Interbank Offering Rate for a 6 months term).

FIGURE 1

Credit rates versus money market rates



(Augmented) Dickey-Fuller tests indicate that all but one interest rates considered here are well described as stationary around a time trend. In what concerns TBA, we could not dismiss that it is not stationary. The t -statistic on the coefficient of the relevant variable in level where the dependent variable is its first difference is displayed in table 2, column 2. Under the null hypothesis of no stationarity this statistic follows a Dickey-Fuller distribution. Five percent critical values for this number of observations are -3.47 or -2.90 , with and without trend, respectively.

TABLE 2

Stationarity tests for interest rates and spreads

| | Interest rates (levels) | | Interest rates (first differences) | Spread (relevant credit rate—TBA) |
|--------------------------|-------------------------|------------|------------------------------------|-----------------------------------|
| | Without trend | With trend | Without trend | Without trend |
| TBA | - 0.50 | - 2.54 | (*) - 6.99 | - |
| 91 to 180 days | - 0.42 | (*) - 4.70 | (*) - 13.39 | (*) - 4.70 |
| 181 days to 1 year | - 1.02 | (*) - 4.16 | (*) - 14.96 | (*) - 3.98 |
| 2 to 5 years | - 0.65 | (*) - 4.89 | (*) - 14.96 | (*) - 3.51 |
| More than 5 years | - 1.41 | (*) - 6.34 | (*) - 14.35 | (*) - 5.70 |

(*) Significant at 5 percent or less.

The same type of tests allowed us to reject that the first differences are not stationary (table 2, column 3). According to the results of ADF tests, all interest rates but TBA would be classified as stationary around a (downward) time

trend. The time trend reflects the fall in inflation observed in the sample years. As it can not be possible that interest rates will be an ever decreasing time series, it is sensible to expect that the fall in interest rates will stop sooner or later, once inflation stabilises. Therefore, we assume the time trend reflects a small sample bias and classify all rates as $I(1)$ (7).

The difference between the credit rates and the money market rate is called the spread. We have tested the hypothesis that the four spreads we have (there is one spread for each of the credit rates) are $I(0)$, or stationary. We have done that again using the Dickey-Fuller test for stationarity. It is clear by the results in table 2 that the no stationarity null hypothesis is comfortably rejected, without any time trend. The spread displays a constant mean through time, so that TBA is cointegrated with each of the credit rates, with cointegrating vector $(1, -1)$. Therefore, we conclude that changes in TBA reflect themselves into *one to one* changes in credit rates.

Having concluded that spreads are stationary, we now turn into the estimation of the speed of adjustment of credit rates to changes in money market rates. We want to answer the following question: How long does the spread take to revert to the mean after a change in TBA? The longer the period the stickier credit rates are, once adjustments in the spread after a change in TBA come through changes in credit rates.

To actually estimate the speed of adjustment we have estimated the following equation for all but one of the four time series of spreads (8):

$$spread_t = b_0 + b_1 \Delta TBA_t + b_2 spread_{t-1} + b_3 spread_{t-2} + u_t \quad (1)$$

where the number of lags is such that residuals show no autocorrelation (9).

Equation (1) is equivalent to the following one, where i is the credit rate (10):

$$\Delta i_t = b_0 + (1 + b_1) \Delta TBA_t - b_3 \Delta spread_{t-1} - (1 - b_2 - b_3) spread_{t-1} + u_t \quad (2)$$

Recalling that $spread = i - TBA$, equation (2) is a reparameterisation of equation (1) in the error correction form. Table 3 presents the estimated coefficients of equation (2).

(7) We wish to thank an anonymous referee who called our attention to this point.

(8) A time trend never turned to be significant.

(9) Only one lag was needed when modelling the 2 to 5 year spread.

(10) We thank an anonymous referee who suggested this formulation.

TABLE 3

Credit rate specification (*t*-statistics between parenthesis)

| Spread (dependent variable) | Constant | Δ TBA | Δ spread (-1) | spread (-1) |
|-----------------------------|-----------------|-----------------|----------------------|---------------------|
| 91 to 180 days | 0.642 (1.75) | 0.181 (1.62) | - 0.219 (- 2.50) | - 0.187 (- 2.20) |
| 181 days to 1 year | 0.719 (2.17) | 0.258 (2.13) | - 0.260 (- 2.83) | - 0.190 (- 2.78) |
| 2 to 5 years | 1.371 (3.40) | 0.198 (1.25) | - | - 0.301 (- 3.85) |
| More than 5 years | 1.478 (3.79) | 0.097 (0.47) | - 0.190 (- 1.88) | - 0.462 (- 4.32) |

The estimated equations allowed us to provide an estimate of the adjustment speed. We computed the time adjustment of the credit rate after a sustained one point increase in TBA. The adjustment in month 0 is given by the coefficient in Δ TBA. This coefficient is in all cases smaller than one, an indication that credit rates already adjust a little in the same month TBA changes. Changes in the following months are implied by the autocorrelation structure of the credit rate equation, and are presented in table 4.

TABLE 4

Cumulative adjustment (in basis points) of credit rate after an increase of 100 basis points in TBA

| Credit rate \ Month | 0 | 1 | 2 | 3 | 4 |
|--------------------------|----|----|----|----|----|
| 91 to 180 days | 18 | 51 | 53 | 61 | 67 |
| 181 days to 1 year | 26 | 59 | 58 | 66 | 71 |
| 2 to 5 years | 20 | 44 | 61 | 73 | 81 |
| More than 5 years | 10 | 69 | 72 | 84 | 89 |

One can infer that following a one point increase in TBA, the average credit rate for lending operations from 91 to 180 days increased by only 67 basis points after four months (or 71 basis points, considering operations between 181 days and 1 year).

We conclude that credit rates do not adjust immediately to changes in money market conditions. There is some degree of stickiness in credit rates, which is not against the credit rationing hypothesis. Nevertheless, as time goes by credit rates do adjust fully to changes in money market rates.

4 — Causality tests

Another possible method of testing for credit rationing is to evaluate the existence of a direct effect of money on loans. In the absence of credit rationing credit should always be on the credit demand schedule, and therefore there should not be any direct influence of a money aggregate. One should only expect to find evidence for the interest rate channel, that is, money would have a direct effect on loan rates and loan rates would also be useful to explain credit. Our objective is therefore to see if there is direct causality from money to credit or, in other words, if the credit rationing channel is operative. Moreover, if credit rationing is not relevant, one would not expect credit to affect economic activity, once interest rates have been taken into account. In other words, empirical evidence favouring a credit channel in the monetary transmission mechanism would add to the conviction that credit rationing is at work.

Using monthly data from January 1990 to October 1997 we proceeded to estimate a multivariate VAR model including four variables: a monetary aggregate ($M1^*$), consistent with a narrow money supply definition; credit to non-financial enterprises and private individuals; the loans and advances rate for operations from 91 to 180 days and economic activity proxied by the Industrial Production Index (IPI) ⁽¹¹⁾.

We have used the logarithms of real money and real credit aggregates, as well as the real loan rate. As one could expect from monthly data of this kind, real money and credit and IPI include a strong seasonal pattern. We have taken the seasonal differences of these variables. As seasonal differencing was not enough to achieve stationarity, we have taken the first difference of the ensuing series. In practice, this means we have used the monthly changes of year-on-year growth rates for these variables ⁽¹²⁾.

The real interest rate was not found to be stationary in levels. The first difference was, so that this variable was considered as a $I(1)$ variable without any significant seasonal pattern. Consequently, monthly changes were used in

⁽¹¹⁾ We did not use monetary base instead of $M1^*$ because there was a significant change in monetary policy during the period under analysis: the reduction of the legal reserve coefficient from 17 per cent to 2 per cent in 1994. This resulted in a substantial one time decline (a 61.2 per cent decrease) of monetary base on November 1994.

⁽¹²⁾ In previous empirical work, we have used comparable but seasonally adjusted series. See Afonso and St. Aubyn (1997).

the VAR. In all cases, stationarity was tested using the Augmented Dickey Fuller test. Table 5 summarises our stationarity tests:

TABLE 5
Stationarity tests for money, credit, loan rate and IPI⁽¹³⁾

| | Level | Seasonal difference | First difference of seasonal difference (a) |
|----------------------|-------------|---------------------|--|
| Real money | — | — 1,38 (12) | (**) — 10,23 (0) |
| Real credit | — | 0,44 (12) | (**) — 7,41 (0) |
| IPI | — | — 1,51 (11) | (**) — 6,46 (4) |
| Real loan rate | — 1,32 (11) | — | (**) — 11,1 (0) |

(a) Except for the interest rate, where seasonal differencing was not necessary.

(**) Significant at 1 percent or less.

The VAR specification is given in equations (3) to (6), with all variables transformed as described above⁽¹⁴⁾:

$$M_t = \text{constant} + \sum_{i=1}^5 \alpha_{mi} M_{t-i} + \sum_{i=1}^5 \beta_{mi} C_{t-i} + \sum_{i=1}^5 \gamma_{mi} R_{t-i} + \sum_{i=1}^5 \delta_{mi} A_{t-i} + u_{mt} \quad (3)$$

$$C_t = \text{constant} + \sum_{i=1}^5 \alpha_{ci} M_{t-i} + \sum_{i=1}^5 \beta_{ci} C_{t-i} + \sum_{i=1}^5 \gamma_{ci} R_{t-i} + \sum_{i=1}^5 \delta_{ci} A_{t-i} + u_{ct} \quad (4)$$

$$R_t = \text{constant} + \sum_{i=1}^5 \alpha_{ri} M_{t-i} + \sum_{i=1}^5 \beta_{ri} C_{t-i} + \sum_{i=1}^5 \gamma_{ri} R_{t-i} + \sum_{i=1}^5 \delta_{ri} A_{t-i} + u_{rt} \quad (5)$$

$$A_t = \text{constant} + \sum_{i=1}^5 \alpha_{ai} M_{t-i} + \sum_{i=1}^5 \beta_{ai} C_{t-i} + \sum_{i=1}^5 \gamma_{ai} R_{t-i} + \sum_{i=1}^5 \delta_{ai} A_{t-i} + u_{at} \quad (6)$$

where:

M — real money ($M1$);

C — real credit (loans and advances),

R — real loan rate $= (1+\text{rate})/(1+p)-1$;

A — IPI index;

¹³ Values within parenthesis denote the number of lags for the dependent variable used in the Dickey-Fuller or Augmented Dickey-Fuller tests.

¹⁴ There was no evidence of cointegration among the $I(1)$ variables, according to Johansen procedure tests, so we proceeded to estimate a VAR using differenced variables without any error-correction terms.

rate — nominal loan rate;
 p — year on year inflation rate.

The main tests we computed are summarised in figure 2 and the F statistics are reported on table 6.

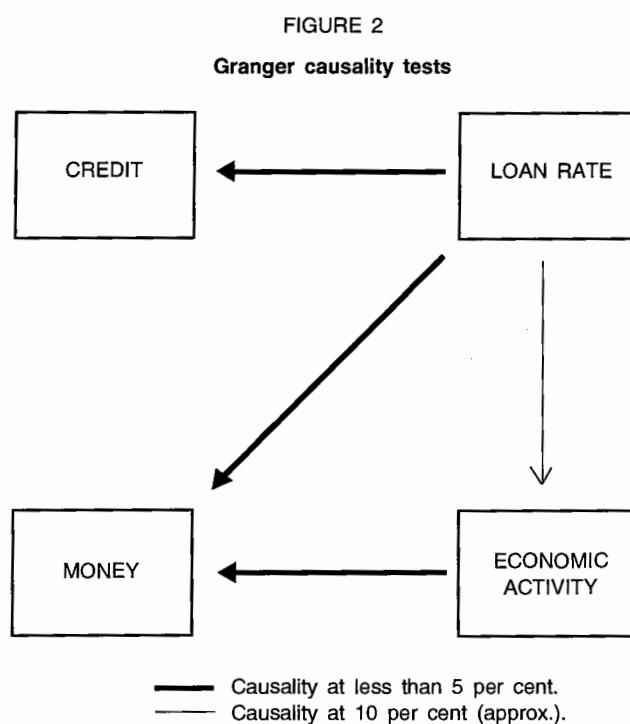


TABLE 6
Causality tests results (F -statistics)

| Dependent variable Causative variable | M1 | Credit | Loan rate | Ec. activity |
|--|-------|--------|-----------|--------------|
| M1 | 0.004 | 0.788 | 0.301 | 0.140 |
| Credit | 0.367 | 0.603 | 0.603 | 0.667 |
| Loan rate | 0.018 | 0.021 | 0.641 | 0.106 |
| Ec. activity | 0.011 | 0.230 | 0.449 | 0.000 |

Note. — Shadowed values imply the existence of causality from the exogenous variable to the dependent variable at less than 5 per cent level.

From the causality tests one can draw some interesting conclusions. There is no evidence of credit rationing in so far as credit does not seem to be influenced by past values of money and is clearly influenced by the interest rate. This can be interpreted as a sign of credit being on the demand curve. Economic activity is not Granger-caused by money or credit once the interest rate is taken into account. Apparently, the interest rate channel is at work and not a credit one. The interest rate is not clearly significant in explaining economic activity, though. Nevertheless, its low p-value and the fact that a rise in the interest rate depresses activity, according to the impulse response function discussed below are in accordance with this interpretation.

The fact that the loan rate causes money comes as no surprise. We have seen previously that loan rates are closely related to money rates, which one would expect to relate to a narrow money aggregate. Also, loan rates cause credit, and credit is related to deposits (a part of M1) by the banks balance sheets. We interpret the remaining causality relationship (from activity to money) as a standard «demand for money» result.

We have computed impulse response functions from the estimated VAR using the Choleski decomposition method¹⁵. Basically, we had to assume:

- i) That innovations to each variable are orthogonal;
- ii) A pattern of contemporaneous effects within variables.

Assumption ii) is closely related to the usually called «ordering» of the VAR. In the impulse response functions below, we have assumed the following order: activity, credit, loan rate, money. This implies that a variable that is positioned before another is not contemporaneously affected by it. For example, activity was assumed not to be contemporaneously affected by any of the other variables, and money to be contemporaneously affected by all other variables. Although there is a degree of arbitrariness in this ordering, it respects the rule that variable *X* is not contemporaneously affected by *Y* if *Y* does not Granger cause *X* at less than the 5 percent significance level. The shape of the impulse response functions was not very different when other orderings were considered.

Impulse response functions concerning the four more significant causal relationships are represented in figures 3 to 6. They trace the response through time of a variable following a standardised change in another variable.

¹⁵ Charemza and Deadman (1997), Enders (1995) and Sarte (1997) discuss this and other methods of identifying impulse response functions.

FIGURE 3

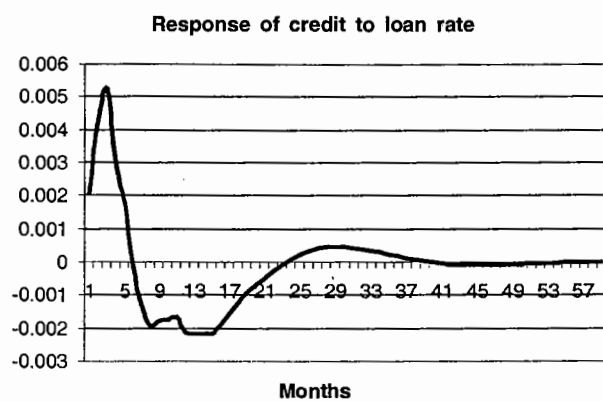


FIGURE 4

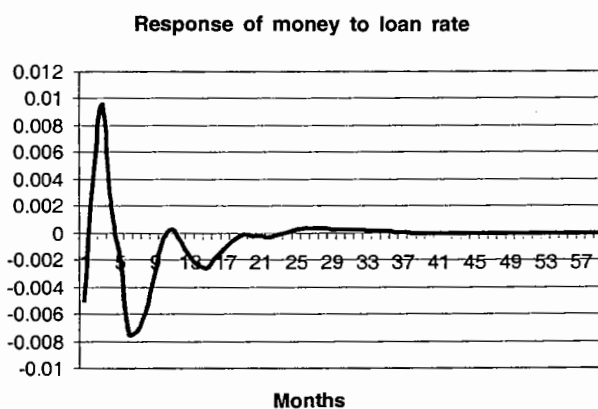


FIGURE 5

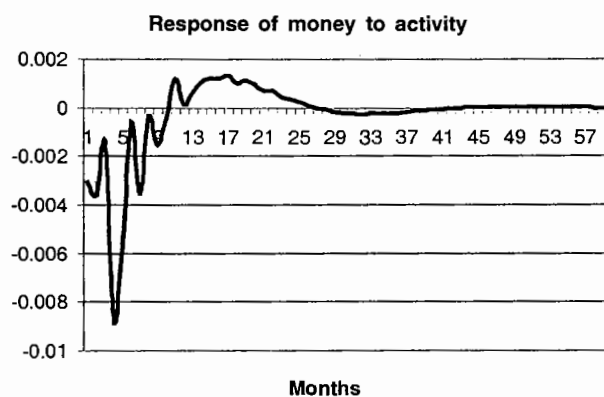
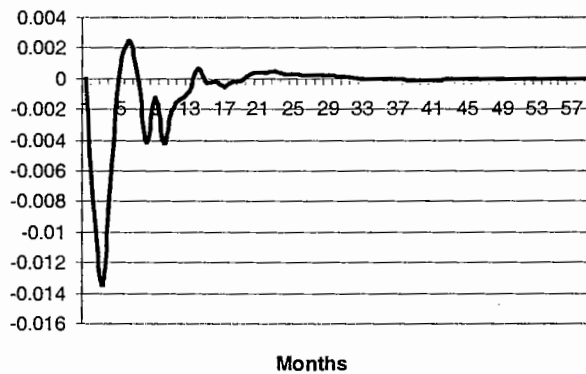


FIGURE 6

Response of activity to loan rate



As one could expect, an increase in the loan rate depresses credit in the mid-term (figure 3), according to the idea that credit is on the demand schedule. The same increase in the loan rate makes money decrease (figure 4). This is probably related to the decrease in credit and therefore in deposits and also to the monetary contraction following a rise in money market interest rates, to which loan rates are closely related. Figure 5 seems to convey the message that money responds positively in the mid-term to activity. Finally, figure 6 implies that a rise in the loan rate depresses economic activity (the interest rate channel), with the caveat that this causal relationship is significant at the 10 per cent level only.

5 — Conclusion

Empirical evidence presented in this paper is not favourable to the existence of credit rationing in Portuguese banking.

Stickiness tests tell us that money market rate changes are transmitted to credit rates changes on a one to one basis. The credit rationing hypothesis would imply that credit rates would observe some degree of independence from money market rates. This is somehow denied by our stickiness results. It is still the case that credit rates take some time to adjust and that some degree of rationing could still be occurring. Clearly there is scope here for further and more detailed empirical research.

The VAR based tests were not more favourable to rationing in credit markets. Narrowly defined money does not seem to condition the level of credit, basically determined by the prevailing interest rate and taking economic activity into account. Also, interest rates are more important for economic activity than

credit or money. The authors are aware that these results could arise from misspecification of credit demand or from the choice of variables (namely for money and economic activity). Results should therefore be interpreted with care and more research on the matter is under way.

ANNEX

Data sources

Loan rates: loan (to non-financial enterprises and private individuals) rates for various loan lengths, Banco de Portugal.

TBA: average rate of Treasury bills auctions, Banco de Portugal.

M1: currency in circulation + demand deposits and other monetary liabilities, Banco de Portugal.

Credit: credit to non-financial enterprises and private individuals, Banco de Portugal.

Economic activity: Industrial production index (1990=100), mainland, adjusted for working days, Instituto Nacional de Estatística.

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